Journal of Agribusiness 24,1 (Spring 2006):79–92

Long-Run Relationship of U.S. and Argentine Maize Prices

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This paper examines the relationship between the maize export prices of the United States and Argentina. The results suggest an asymmetric nature of price adjustment. This could be due to the fact that the maize market is characterized by significant concentration. The larger market share of exports by the United States reflects the influences on export price dynamics. The structure of maize trade is such that U.S. markets are largely insulated from influences flowing from Argentina, while Argentine maize prices are not insulated from U.S. influences.

Key Words: cointegration, maize, price dynamics, threshold adjustment

International price relationships in agricultural commodity markets have been of considerable interest to economists because of issues such as price transmission, market integration, and price leadership. If an international commodity market were to be integrated, one would expect, through arbitrage and/or substitution, the prices charged by major exporters of that commodity to move together. A price change in one market will be followed by a gradual similar price change in the other market, thereby disallowing the prices of the commodity to diverge too far from each other. When a group of prices move proportionally to one another over time, then the Law of One Price (LOP) holds. From an econometric point of view, this would imply that prices of the commodity should be cointegrated.

International trade in many products is dominated by relatively few agents on either or both sides of the market (Karp and McCalla, 1983). The international maize market is a case in point. The major maize exporting countries over the past decade include the United States (U.S.) and Argentina. The U.S. is the world's largest maize exporter, accounting for approximately 70% of total maize exports in the world maize market. In comparison, Argentina's exports of maize over the last decade have been considerably lower, accounting for approximately 13% of total global exports [U.S. Department of Agriculture/Foreign Agricultural Service (USDA/FAS), 2001].

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Maize is one of the world's most important grains, and is produced and consumed across the globe. Maize provides the basic staple for human consumption in much of Sub-Saharan Africa, Central America, the Andean region, and parts of Asia. It is also used as feed for livestock in middle- and upper-income countries (Falcon and Naylor, 1998). During the last decade, the total quantity of maize produced globally averaged approximately 560 million tons. Over this same period, the total quantity of maize traded averaged 65 million tons, which is approximately one-eighth of total world production (Meng and Ekboir, 2001).

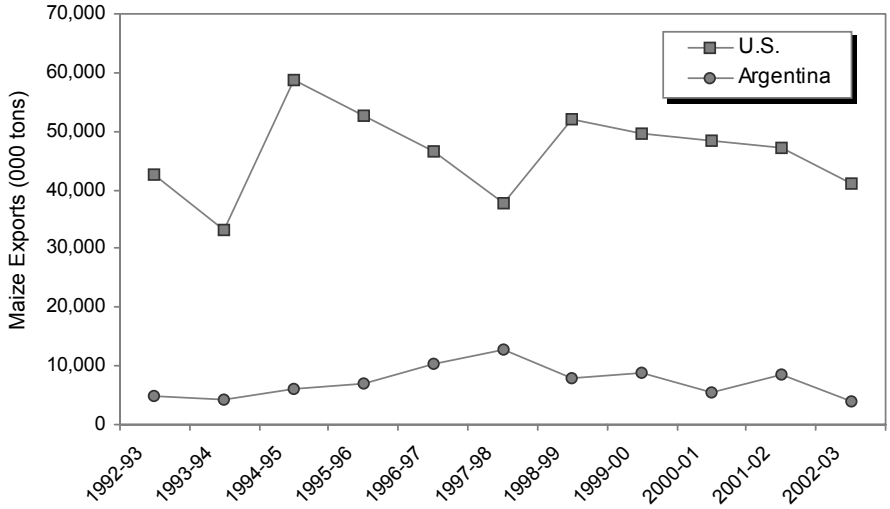
The structure of the world maize market can be characterized as an oligopoly (Ray, 2003). This perception is driven by the fact that the U.S. and Argentina are the two major exporters of maize and are geographically widely separated. Together, the U.S. and Argentina account for over 80% of the world export of maize [Meng and Ekboir (2001), as developed from USDA/FAS 2001 data]. Traditionally, the maize export market has been dominated by the U.S., whereas Argentina by comparison exports significantly lower volumes of maize (Lence, 2000). China is somewhat of an anomaly, acting as both a major exporter and importer during the last decade (Meng and Ekboir, 2001).[1] As the dominant exporter, the U.S. would be expected to act as a price leader and Argentina as a price follower.

If the market system were perfectly competitive, then price increases should be transmitted to the same extent as price decreases (Goletti and Babu, 1994). However, because the international maize export market exhibits a structure more consistent with an oligopoly than a perfectly competitive market, the price transmission is likely to be asymmetric. An important point frequently overlooked in the literature is that if the underlying process of adjustment is asymmetric, the tests for cointegration and its extensions are misspecified (Enders and Siklos, 2001). Recent developments in time-series analysis have recognized the potential for threshold type adjustments in error correction models. Threshold effects occur when price increases bring a different response than price decreases.

Based on a review of the literature, no previous study has examined the price dynamics in the international maize market. Argentina's current and future export potential could have a major impact on U.S. maize exports. Years of fiscal reform and improvements in infrastructure have laid the groundwork for Argentina's emergence as a serious year-round, global competitor to the U.S. in the maize market. The extent of such competition is reflected in figure 1, which illustrates the volume of maize exports for the U.S. and Argentina over the 1992–93 through 2002–03 study period. Despite the difference in market share for maize exports, it can be seen from figure 1 that U.S. maize exports generally tend to move in an opposite direction to those of Argentina.

Argentina has been expanding into the Middle East and North African markets, where the U.S. had already established a strong presence (USDA/FAS, 2003b). Further, when the European Union implemented a ban on nonapproved genetically

[1] During the period analyzed, China has emerged as a significant net exporter of maize for the last three marketing years (King, 2001; Ray, 2003).



Source: Grain: World Markets and Trade, USDA/FAS.

Figure 1. U.S. and Argentine maize exports (000s of tons), 1992–93 through 2002–03

modified (GM) maize in 1997–98, U.S. exports of maize plummeted (figure 1). Because Argentina planted only the approved varieties of maize, its exports expanded—driven primarily by demand from Spain, Portugal, and the UK. The intense competition between U.S. and Argentina for maize exports over the last decade warrants serious consideration in investigating the long-run relationship between Argentina and U.S. export prices of maize.

The primary objective of this paper is to examine the relationship between the maize export prices of the U.S. and Argentina. Cointegration analysis is a useful technique for investigating the price relationships in international commodity markets (see Ardeni, 1989; Baffes, 1991; Goodwin and Schroeder, 1991; Asche, Bremnes, and Wessells, 1999; Mohanty, Meyers, and Smith, 1999; Bessler, Yang, and Wongcharupan, 2002).[2] The motivation of such research is to define market boundaries and identify any causal relationships that may exist among prices for commodities in the same market. If the commodities belong to the same market, there will exist a long-run relation between prices driven by arbitrage and/or substitution. In this way, by examining the relationships between prices over time, the relevant market can be defined. The extent of the market (Stigler and Sherwin,

[2] Ardeni (1989) attempted to verify the Law of One Price (LOP) focusing on a small number of commodities which included wheat in addition to wool, beef, sugar, tin, tea, and zinc. Baffes (1991) conducted a similar study to test whether the LOP holds for wheat as well as beef, tea, sugar, wool, zinc, and tin. Both studies employed cointegration tests. Goodwin and Schroeder (1991); Mohanty, Meyers, and Smith (1999); and Bessler, Yang, and Wongcharupan (2002) examined the price relationships in the international wheat market following a cointegration approach. Asche, Bremnes, and Wessells (1999) used cointegration to test for price relationships in the world salmon market.

1985) implies that the commodity prices in different geographical locations tend to move together.

Given the imperfect market structure of the global maize market, this study aims to test for the presence of cointegration in the presence of asymmetric error correction. An attractive method of estimation and testing follows the approach reported by Enders and Siklos (2001). Noting there is a correspondence between error correction models representing cointegrating relationships and autoregressive models of an error term, the approach used here is an application of the Enders and Siklos (2001) threshold autoregressive (TAR) and momentum-threshold autoregressive (M-TAR) method of adjustment.

Under the TAR model, the underlying price series might exhibit asymmetry where one price remains above the other price, but for short intervals, the lower price peaks above the higher price (Ghoshray, 2002). For the M-TAR model, the underlying price series might display asymmetry of the following form: When prices are decreasing, the gap between the prices decreases at a faster rate as opposed to the case when both prices are increasing (Ghoshray, 2002). In both cases, this type of asymmetric price behavior differs from the symmetric adjustment scenario where both prices move together and any transitory deviation in "levels" or "rates of change" is corrected in a symmetric manner.

A further objective of this paper is to use these price relationships to provide information on the existence of price leadership. In particular, we address the question of whether the U.S. is driving the maize export prices. The issue of price leadership is of interest because the maize market is perceived to be imperfectly competitive, and as the dominant exporter, the U.S. is capable of exerting market power—a prerequisite for any kind of price leadership role. Previous studies (e.g., Smith, Goodwin, and Holt, 1995; Asche, Bremnes, and Wessells,1999; Mohanty, Meyers, and Smith,1999; Bessler, Yang, and Wongcharupan, 2002)[3] have tested for price leadership in various international commodity markets.

The remaining sections of the paper are organized as follows. First, the econometric model is presented, followed by a description of the data. The empirical results are then discussed, and in the final section conclusions are drawn from the analysis.

The Econometric Model

The Engle and Granger (1987) two-step method to test for cointegration between two variables, say prices P_{1t} and P_{2t}, entails using ordinary least squares (OLS) to estimate the long-run relation of the two prices. This is represented by:

$$P_{1t} = \alpha + \beta P_{2t} + g_t . \tag{1}$$

[3] Smith, Goodwin, and Holt (1995) and Mohanty, Meyers, and Smith (1999) employed Granger causality to test for price leadership in the international wheat market. Asche, Bremnes, and Wessells (1999) and Bessler, Yang, and Wongcharupan (2002) used weak exogeneity tests to determine whether price leadership exists in the world salmon and wheat markets, respectively.

Here, P_{1t} and P_{2t} are nonstationary I(1) prices; α and β are the estimated parameters, where α is an arbitrary constant that accounts for the differential (transfer costs and quality differences), and β denotes the price transmission elasticity; and g_t is the error term which may be serially correlated. To test the hypothesis, equation (1) is estimated using OLS. The second step advocates a Dickey-Fuller test on the estimated residuals of (1), given by:

$$\Delta g_t = \gamma g_{t-1} + \omega_t, \tag{2}$$

where ω_t is a white noise error term. If ω_t is not white noise, an augmented Dickey-Fuller (ADF) test may be used where lagged values of Δg_t may be added to (2). Rejecting the null hypothesis of no cointegration (H_0: $\gamma = 0$) implies the residuals of (2) are stationary. Thus, (1) is like an attractor such that its pull is strictly proportional to the absolute value of g_t.

However, Enders and Granger (1998) and Enders and Siklos (2001) argue that the tests for cointegration and its extensions are misspecified if adjustment is asymmetric. They consider an alternative specification, termed the threshold autoregressive (TAR) model, whereby (2) can be rewritten as:

$$\Delta g_t = I_t \gamma_1 g_{t-1} + (1 - I_t)\gamma_2 g_{t-1} + \omega_t, \tag{3}$$

where I_t is the Heaviside Indicator function such that:

$$I_t = \begin{cases} 1 & \text{if } g_{t-1} \geq \tau, \\ 0 & \text{if } g_{t-1} < \tau. \end{cases} \tag{4}$$

This specification allows for asymmetric adjustment. If the system is convergent, then the long-run equilibrium value of the sequence is given by $g_t = \tau$. The sufficient conditions for the stationarity of g_t are $\gamma_1 < 0$, $\gamma_2 < 0$, and $(1 + \gamma_1)(1 + \gamma_2) < 1$ (Petrucelli and Woolford, 1984). In this case, if g_{t-1} is above its long-run equilibrium value, then adjustment is at the rate γ_1; if g_{t-1} is below long-run equilibrium, then adjustment is at the rate γ_2. The adjustment would be symmetric if $\gamma_1 = \gamma_2$. However, if the null hypothesis H_0: $(\gamma_1 = \gamma_2)$ is rejected, then using the TAR model we can capture signs of asymmetry. If, for example, $-1 < \gamma_1 < \gamma_2 < 0$, then the negative phase of the g_t series will tend to be more persistent than the positive phase.

In the above case, it is necessary to estimate the value of the threshold that will be equal to the cointegrating vector. A method of searching for a super-consistent estimate of the threshold was undertaken by using an approach proposed by Chan (1993). To utilize Chan's methodology, the estimated residual series was sorted in ascending order, i.e., $g_1 < g_2 < \ldots < g_T$, where T denotes the number of usable observations. According to this method, the largest and smallest 15% of the g_t series were eliminated, and each of the remaining 70% of the values were considered as possible thresholds. For each of the possible thresholds, the equation was estimated using (3)

and (4). The estimated threshold yielding the lowest residual sum of squares was deemed to be the appropriate estimate of the threshold.

In equation (4), the Heaviside Indicator depends on the level of g_t. Enders and Granger (1998) and Enders and Siklos (2001) suggest an alternative such that the threshold depends on the previous period's change in g_t. The Heaviside Indicator given by (4) can be set to the Momentum-Heaviside Indicator as:

$$I_t = \begin{cases} 1 & \text{if } \Delta g_{t-1} \geq \tau, \\ 0 & \text{if } \Delta g_{t-1} < \tau. \end{cases} \tag{5}$$

In this case, the series g_t exhibits more momentum in one direction than the other. The model given by (3), along with equation (5), depicts the momentum threshold autoregression (M-TAR) model. The M-TAR model can be used to capture a different type of asymmetry. If, for example, $|\gamma_1| < |\gamma_2|$, the M-TAR model exhibits little adjustment for positive Δg_{t-1}, but substantial decay for negative Δg_{t-1}. In other words, increases tend to persist, but decreases tend to revert quickly back to the attractor irrespective of where disequilibrium is relative to the attractor.

To estimate the threshold, Chan's (1993) methodology is employed. The estimated residual series was sorted in ascending order: $\Delta g_1 < \Delta g_2 < \ldots < \Delta g_T$, where T denotes the number of usable observations. As before with the TAR model, the largest and smallest 15% of the g_t series were eliminated, and each of the remaining 70% of the values were considered as possible thresholds. For each of the possible thresholds, the equation was estimated using (3) and (5). Again, the estimated threshold yielding the lowest residual sum of squares was deemed to be the appropriate estimate of the threshold.

To implement in this test the case of the TAR or M-TAR adjustment, the Heaviside Indicator function is set according to equation (4) or equation (5), respectively, and equation (3) is estimated accordingly. The Φ-statistic for H_0: ($\gamma_1 = \gamma_2 = 0$), the null hypothesis of stationarity of g_t, is recorded. The value of the Φ-statistic is compared to the critical values computed by Enders and Siklos (2001). If the null hypothesis can be rejected, it is possible to test for asymmetric adjustment since γ_1 and γ_2 converge to multivariate normal distributions (Tong, 1990). The F-statistic is used to test for the null hypothesis of symmetric adjustment, H_0: ($\gamma_1 = \gamma_2$). Diagnostic checking of the residuals is undertaken using the Ljung-Box Q-tests to ascertain whether the ω_t series is a white noise process. If the residuals are correlated, equation (3) needs to be reestimated in the form:

$$\Delta g_t = I_t \gamma_1 g_{t-1} + (1 - I_t)\gamma_2 g_{t-1} + \sum_{i=1}^{p} \phi_i \Delta g_{t-i} + \omega_t. \tag{6}$$

Equation (6) is comparable to (3) except that it incorporates lagged first differences of the dependent variable to correct for the autocorrelation in the error term ω_t. The Schwarz Bayesian Criterion (SBC) is used to determine the lag length.

Empirical Analysis

Data

The data used for this analysis are monthly average export price quotations (FOB) from July 1993 through January 2003. The maize prices considered in this study are for the Argentine Rosario and U.S. Yellow No. 3 from the Gulf port. The data were obtained from *World Grain Statistics* published by the International Grains Council. All prices are quoted in U.S. dollars. The subsequent analysis of the data is carried out on the logarithm of prices. Figure 2 illustrates the maize export prices from the U.S. and Argentina. As observed from this graph, the prices appear to move together over time.

Empirical Results

The prices were initially tested for their order of integration. Table 1 presents the results of the unit root tests for each of the price series, expressed in log levels. Based on the ADF test results for the variables in levels and growth form, the prices are nonstationary I(1).

Following the Engle and Granger (1987) methodology, the first step entails estimating the long-run equilibrium relationship given by equation (1) and then conducting the ADF test on the residuals of (1). The results of the test are reported in table 2. The ADF test was conducted with 2 lags determined by the Schwarz Bayesian Criterion (SBC). The key point to note is that the ADF t-statistic is !6.04, indicating the null of no cointegration can be rejected, and therefore both the prices are cointegrated.

The residuals of (1) are estimated in the form of the TAR and M-TAR models. The consistent estimate of the threshold value using Chan's (1993) method was found to be !0.071 and !0.039, respectively. The results of the TAR model are shown in the second numeric column of table 2. The point estimates are calculated as γ_1 ' !0.46 and γ_2 ' !0.30, which have the correct signs for convergence. The statistic Φ ' 19.41 is greater than the 5% critical value, implying the null hypothesis of no cointegration can be rejected. Given that cointegration is detected, the null hypothesis of symmetric adjustment can be tested using the standard F-distribution. The sample value of F ' 2.00 has a p-value of 0.16. Consequently, the null of symmetric adjustment cannot be rejected.

The last column of table 2 reports the results using a consistent M-TAR model. The point estimates of γ_1 ' !0.33 and γ_2 ' !0.76 suggest convergence. The statistic Φ ' 24.56 allows us to reject the null hypothesis of no cointegration at the 5% significance level. The null hypothesis of symmetric adjustment provides a sample value of F ' 9.72 with a p-value of 0.00, implying the null can be rejected. Given the finding of asymmetric adjustment, the power of the Φ-statistic in this case exceeds that of the Dickey-Fuller test (Enders, 2001). Finally, the Ljung-Box Q-statistic reveals that none of the models suffer from problems of serial correlation.

Source: World Grain Statistics, International Grains Council.

Figure 2. U.S. and Argentine maize export prices (U.S. $/ton), July 1993 through January 2003

The estimates of γ_1 and γ_2 are expected to be negative, suggesting convergence for the M-TAR model. Since $|\gamma_1| < |\gamma_2|$, the M-TAR model exhibits little adjustment for positive Δg_{t-1}, but substantial decay for negative Δg_{t-1}. Specifically, increases tend to persist, but decreases tend to revert quickly back to the attractor. Thus, when using the M-TAR consistent model, the results suggest we can find evidence of asymmetric adjustment.

The positive finding of cointegration with threshold adjustment justifies the estimation of the following threshold error correction model (TECM) of the form:

$$\Delta P_t^{AR} = \mu_1 ECM_{t-1}^{+} + \mu_2 ECM_{t-1}^{-} + \sum_{i=1}^{p} \psi_i \Delta P_{t-i}^{AR} + \sum_{i=1}^{p} \delta_i \Delta P_{t-i}^{US} + u_t^{AR}, \tag{7}$$

$$\Delta P_t^{US} = \pi_1 ECM_{t-1}^{+} + \pi_2 ECM_{t-1}^{-} + \sum_{i=1}^{p} \eta_i \Delta P_{t-i}^{AR} + \sum_{i=1}^{p} \lambda_i \Delta P_{t-i}^{US} + u_t^{US}, \tag{8}$$

where P^{AR} and P^{US} denote the export prices of Argentina and the U.S., respectively; *ECM* refers to the error correction term; and both u_t^{AR} and u_t^{US} are white noise errors. The number of lags (p) is determined using the SBC. The results from the TECM are reported in table 3.

The p-values for the error correction terms indicate that P^{AR} adjusts to deviations from the long-run equilibrium at the 10% significance level. In the short run there is no evidence of any causality. The lagged changes in both the prices affect movements in the short run. However, there is no indication of Granger causality. Finally, the diagnostics suggest the model does not suffer from serial correlation.

Table 1. Results of the Augmented Dickey-Fuller Test

	ADF *t*-Test	
Price Series	Levels	Growth Form
Argentina	!2.50 [1]	!7.10* [0]
United States	!2.11 [1]	!7.20* [0]

Notes: An asterisk (*) denotes significance at the 5% level. Numbers within brackets denote lag length determined by the Schwarz Bayesian Criterion.

Table 2. Estimates of the Long-Run Price Relationship

	Engle-Granger		TAR		M-TAR	
Coefficient/Statistic	Estimate	(*t*-Statistic) [*p*-Value]	Estimate	(*t*-Statistic) [*p*-Value]	Estimate	(*t*-Statistic) [*p*-Value]
γ_1	!0.39	(!6.04)	!0.46	(!5.72)	!0.33	(!5.03)
γ_2	NA		!0.30	(!3.50)	!0.76	(!5.68)
ϕ_1	0.37	(4.60)	0.38	(4.70)	0.34	(4.33)
ϕ_2	0.07	(0.85)	0.06	(0.73)	0.14	(1.56)
Φ-Statistic	NA		19.41*		24.56*	
H_0 (γ_1 ' γ_2)	NA		2.00	[0.16]	9.72	[0.00]
Ljung-Box *Q*-Statistic	0.73	[0.94]	0.53	[0.97]	1.64	[0.80]

Notes: An asterisk (*) denotes significance at the 5% level. The values corresponding to Φ are compared with the Φ tables computed by Enders and Siklos (2001).

Table 3. Estimates of the Threshold Error Correction Model (TECM)

	ΔP_t^{AR}		ΔP_t^{US}	
Coefficient/Statistic	Estimate	[*p*-Value]	Estimate	[*p*-Value]
$ECM_{t\&1}^{\%}$	!0.14	[0.07]	0.13	[0.08]
$ECM_{t\&1}^{\&}$	!0.52	[0.00]	0.16	[0.27]
$\Delta P_{t\&1}^{AR}$	0.49	[0.00]	0.03	[0.76]
$\Delta P_{t\&1}^{US}$	0.03	[0.79]	0.34	[0.00]
Durbin-Watson (DW) Statistic	1.97		1.96	
Ljung-Box *Q*-Statistic	1.67	[0.79]	2.42	[0.65]

Examining the results of the weak exogeneity test, the *F*-statistic for the null hypothesis that the Argentinean price is weakly exogenous [i.e., H_0: (μ_1 ' μ_2 ' 0)] is estimated to be 7.11 with a *p*-value of 0.00, implying rejection of the null hypothesis. Testing the null hypothesis that the U.S. price is weakly exogenous [H_0: (π_1 ' π_2 ' 0)], the *F*-statistic is estimated to be 2.09 with a *p*-value of 0.12. The significance level of the U.S. price (12%) does not allow us to reject the null hypothesis that the U.S. price is weakly exogenous.

Discussion of Results

Using the Engle-Granger framework, a long-run relationship is found to exist between the U.S. and Argentine maize prices. The results of the cointegration test imply that the prices for both countries are integrated. This result supports the notion that in the world maize market, better quality and more effective transmission of market information have led to more stability in export pricing, with Argentina's export prices being closely linked to U.S. prices (USDA/FAS, 1998).

When testing for asymmetric adjustment, we find there is little evidence when using the consistent TAR model. However, where evidence of asymmetry is detected, the consistent M-TAR model best fits the data. Using the M-TAR model, the absolute value of γ_1 is estimated to be less than the absolute value of γ_2. This finding indicates the M-TAR model exhibits slow amounts of adjustment toward equilibrium for a positive change in the equilibrium relationship, but a substantial amount of adjustment for a negative change in the equilibrium relationship. Thus, positive changes tend to persist, while negative changes tend to revert quickly back to the attractor. The results of the M-TAR consistent model appear to suggest that the underlying process of adjustment is asymmetric.

The path of adjustment to equilibrium exhibiting M-TAR type adjustment to the long-run equilibrium relation is relatively faster when the price differential is decreasing as opposed to the case when it is increasing. When prices are decreasing, the gap between the prices decreases at a faster rate in contrast to the case when both prices are increasing. The finding of this asymmetric pattern of adjustment supports the contention that the international maize market is characterized by imperfect competition. As stated earlier, imperfect competition gives rise to asymmetric price transmission. When the Argentine maize price is increasing, the U.S. price increases at a faster rate, and when the Argentine maize price is decreasing, the U.S. maize price decreases at a relatively slower rate. Due to Argentina's lower share of total exports of maize, Argentine maize prices would have to decrease at a faster rate in order to retain their market share.

This relative price movement can be observed from the graph in figure 2. Apart from the price increases in 1995–96, there is a general downward trend of the maize export prices for both the U.S. and Argentina. Over the periods when the Argentine maize export price is decreasing, the U.S. maize price decreases at a relatively slower rate. This verifies the inference made from the econometric analysis.

The nature of price dynamics might explain the impact that Argentina could have been making on the international maize market. Since the early 1990s, Argentina has posed a challenge to the U.S. for maize exports by expanding into Middle East and North African markets, where a strong U.S. presence had already been established. The North African market—comprised of Egypt, Algeria, Tunisia, and Morocco—imports a total of 8.6 million tons of maize. Egypt and Morocco have traditionally purchased 80% of their maize from the U.S., with Argentina supplying the rest. However, for the markets of Algeria and Tunisia, Argentina has been making inroads (USDA/FAS, 2003b) at the expense of the U.S. Since 1998–99, global trade of maize has expanded by 7 million tons to 75.7 million tons, while U.S. exports have shrunk by 5 million tons (USDA/FAS, 2003a).

One might expect that the differences between U.S. and Argentina maize price dynamics may also reflect the differences in quality of maize exports. A rise in the price of lower quality maize might lead to a faster rise in the price of higher quality maize, while a decline in lower quality maize prices might have little effect because importers might be willing to pay a premium for higher quality maize. It is surprising to find that even though Argentine maize is considered to be of higher quality than the maize exported by the U.S. (U.S. Congress, Office of Technology Assessment, 1989), the dynamic behavior of price adjustment does not support this hypothesis. Rather, the results suggest a more important role is played by market share than by quality differences. Further, the finding of U.S. price leadership can be attributed to the fact that the U.S. is the largest exporter of maize. The price of U.S. maize appears to evolve independently. Argentina's exports of maize are significantly lower, and therefore Argentina seems to act as a price follower. Argentinean prices respond to correct for any deviation caused in the long-run relationship with U.S. maize prices.

Conclusion

A novel method advanced by Enders and Siklos (2001) was employed to test the hypothesis of cointegration with asymmetric adjustment, and was applied to the U.S. and Argentinean maize export prices. Both of these countries are major exporters of maize. The method extends the Engle-Granger procedure allowing for either TAR or momentum TAR adjustment toward the cointegrating vector. If asymmetry exists, the power of the M-TAR test is higher than that of the Engle-Granger test (Enders, 2001).

The results of the cointegration tests under both symmetric and asymmetric adjustment reveal that the maize export prices of the U.S. and Argentina are tied together in a long-run relationship—implying arbitrage from the supply side or substitution from the demand side binds the prices together over time. This result has implications for grain traders and farmers, and supports the contention that in recent years Argentina has emerged as a tough competitor to the U.S. If the degree of integration were to be low, then grain exporting companies could earn excess profits by buying maize from farmers at very low prices.

In testing for asymmetric adjustment, little evidence is found when using the consistent TAR model. However, where evidence of asymmetry is detected, the consistent M-TAR model best fits the data. Price asymmetry is the market's inability to respond similarly to either a rise or a fall in prices. An over-reaction to price increases while remaining opposed to price decreases could indicate the presence of market power among traders of the dominant exporting country. It does not come as a surprise that the findings of this study suggest an asymmetric nature of price adjustment. Given that the international maize market has a structure more consistent with a duopoly than a perfectly competitive market, the price transmission is expected to be asymmetric. The U.S. holds the largest market share in the international maize market which can be cited as the source of asymmetric price transmission.

As noted earlier, the international maize market is highly concentrated. This nature of export market concentration supports the contention that the world maize market is imperfectly competitive. Under these circumstances, some form of strategic interaction may be expected between the major exporters of maize. The results indicate Argentina shows greater responsiveness to decreasing U.S. prices than increasing U.S. prices. Further, inventory-holding behavior by the U.S. and Argentina may also lead to asymmetries, as high international price expectations lead to stock accumulation (Rapsomanikis, Hallam, and Conforti, 2003).

The weak exogeneity tests on the threshold error correction model show that the U.S. is found to act as a price leader. This finding suggests the structure of maize trade is such that the U.S. markets are largely insulated from influences flowing from Argentina, while Argentine maize prices are not insulated from U.S. influences. The result has important implications for grain traders, as it reflects the mechanisms used by each country to implement pricing policies for their exports of maize. Argentine maize traders might be using the U.S. maize price as a reference when setting their own export prices. This is true of exporting firms, as Argentine export prices tend to be more highly correlated with the U.S. prices than domestic prices (Lence, 2000). This result supports the assertion of Perkins, Snickers, and Geldard (1984) that the U.S., with the largest market share, effectively sets the world price, and other exporters adjust their prices in response to any deviation in the U.S. price. The results further suggest that agricultural policies of Argentina which aim toward improving the quality of maize to compete with the U.S. would not have an effect on the long-run relation between the prices.

This study highlights the fact that the maize export prices are highly integrated, and the prices are set through a process of strategic interaction where the U.S. plays the role of a price leader and Argentina responds in an asymmetric fashion to any variations in the U.S. price. These results are of significance as they add to our understanding of how the international maize market operates.

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